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Interregional and International Risk Sharing and Lessons for EMU

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Introduction[*]

The project of European Monetary Union (EMU) never ceases to raise questions about the capacity of country members to cope with shocks that hit them independently of the rest, and repeated efforts are made to draw lessons from similar experiences of regions within countries. One particular protection that regions have but the country members of EMU will not – or only to a negligible extent – has received wide attention: namely, net transfers through the central government. A region with a large central-government budget undergoing an adverse shock may get notable help through public transfers. But there are other, more decentralized mechanisms that operate to attenuate regional shocks. How important are these in relation to public aid? A recent article by Asdrubali, Sørensen and Yosha (1996) suggests an intriguing way of coming up with answers. These authors propose a method of assessing how much smoothing regions get via insurance and credit as opposed to the central budget, and their method requires little more regional data than figures for output, distributed income (before central government net transfers), disposable income (after the transfers), and consumption. The insurance in question comes from the holding of claims against the output of other regions. The credit channel relates to borrowing from other regions. Based on an application to the US, Asdrubali, Sørensen and Yosha (ASY) conclude that insurance is far more important than credit as a source of smoothing of regional shocks in this country. But credit itself is nearly twice as important as net transfers from the central government. Thus, market forces evidently play an enormous role.

These results, and the reasoning from which they stem, are sufficiently important to merit close scrutiny. In this paper, we will probe more deeply into ASY's method and findings, propose major revisions in their framework, and test the model anew on the basis of the same US evidence as theirs, subject to our revisions. We will then extend the tests outside the US, both, to other individual countries and to groups of them, specifically, the OECD and the European Union. Finally, we will try to draw lessons for EMU.

We shall find that ASY's method holds up extremely well in the US under the more demanding conditions we impose. The method will prove successful in Canada too, but it will fail for Italy and the United Kingdom, the

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other two individual countries we shall study. The international application to groups of countries will turn out particularly fruitful: it will yield implications about the impact of moving from national monetary independence to monetary integration. We will conclude that even though the surrender of monetary policy will undoubtedly reduce the theoretical capacity of the members of EMU to smooth shocks via public action (macroeconomic policy), the change will promote smoothing of shocks via market channels. This last conclusion is perhaps important enough to merit presentation as the central message of the paper. But we shall give the method of investigation pride of place.

The order of the discussion will be, first, the general approach, second, the US results, third, the further individual country tests, fourth, the international tests, and fifth, the conclusions.

I. The General Approach

(a) Original Formulation and Support

Suppose we have a panel of data for per capita regional output Y_i (where i stands for the individual region), per capita regional personal income PI_i, per capita regional disposable income DI_i, and per capita regional consumption C_i, all stated in real terms. Let us begin with the identity

$$Y_i = \frac{Y_i}{PI_i}\frac{PI_i}{DI_i}\frac{DI_i}{C_i}C_i \qquad (1)$$

and next take logarithms and first differences, thereby obtaining:

$$\begin{aligned}\Delta \log Y_i &= (\Delta \log Y_i - \Delta \log PI_i) + (\Delta \log PI_i - \Delta \log DI_i) \\ &\quad + (\Delta \log DI_i - \Delta \log C_i) + (\Delta \log C_i) \qquad (2)\end{aligned}$$

If we multiply both sides of equation (2) by $\Delta \log Y_i$, subtract the means of the term on the left and those of the four terms (in separate parentheses) on the right over the study period, and then take expected values, we will have the variance of the change in the log of Y_i ($\Delta\log Y_i$) on the left and the sum of the covariances of this term with $\Delta\log Y_i - \Delta\log PI_i$, $\Delta\log PI_i - \Delta\log DI_i$, $\Delta\log DI_i - \Delta\log C_i$, and $\Delta\log C_i$, respectively, on the right. Finally, if we divide both sides of the last equation by the variance of $\Delta\log Y_i$, we get

$$1 = \beta_K + \beta_G + \beta_C + \beta_U \qquad (3)$$

In equation (3), the β terms correspond to OLS estimates resulting from the following regressions:

$$\begin{aligned}
\Delta \log Y_i - \Delta \log PI_i &= \alpha_K + \beta_K \Delta \log Y_i + \mu_{iK} \\
\Delta \log PI_i - \Delta \log DI_i &= \alpha_G + \beta_G \Delta \log Y_i + \mu_{iG} \\
\Delta \log DI_i - \Delta \log C_i &= \alpha_C + \beta_C \Delta \log Y_i + \mu_{iC} \\
\Delta \log C_i &= \alpha_U + \beta_U \Delta \log Y_i + \mu_{iU}
\end{aligned} \qquad (4)$$

(where β_K is the covariance between $\Delta \log Y_i$ and $\Delta \log Y_i - \Delta \log PI_i$ divided by the variance of $\Delta \log Y_i$, etc.) In actual estimation, the α's could be nil. Suppose, however, we introduce a separate α for each date, and therefore as many of these constants as there are years in the observation period. These constant terms will then capture any common element in the growth rate of regional per capita output at the separate dates. Consequently, the coefficients of the $\Delta \log Y_i$ terms in the regressions should reflect essentially impulses stemming from the regional deviations of output growth from the national growth rates.

Since the sum of these four β coefficients must equal one (identity (3)), it may be impossible to estimate all four coefficients statistically. There are several ways of handling the problem, of which ASY chose that of regressing jointly $\Delta \log PI_i$, $\Delta \log DI_i$, and $\Delta \log C_i$, respectively, instead of $\Delta \log Y_i - \Delta \log PI_i$, $\Delta \log PI_i - \Delta \log DI_i$, $\Delta \log DI_i$, $- \Delta \log C_i$, and $\Delta \log C_i$, on $\Delta \log Y_i$, thereby obtaining three coefficients, λ_K, λ_G, and λ_C, and then subsequently interpreting β_K as $1 - \lambda_K$, β_G as $\lambda_K - \lambda_G$, β_C as $\lambda_G - \lambda_C$ and β_U simply as λ_C. ASY also evidently estimated β_K, β_G, β_C jointly, β_U separately, and found, both, that their estimates conform to $\beta_K + \beta_G + \beta_C + \beta_U = 1$, and that these estimates yield nearly identical values to those gotten with the first method. Quite significantly, ASY also corrected for heteroskedasticity, or the greater variances of the $\Delta \log Y_i$ values for the smaller US states than the larger ones. We shall examine shortly the issues raised by identity (3) and heteroskedasticity. For the moment, our attention centers on ASY's interpretation of the β coefficients. Quite critically, they propose viewing β_K as a measure of the smoothing of regional shocks to per capita output resulting from cross-regional ownership of claims to output, β_G as a measure of the smoothing of these shocks through the central government budget, β_C as a measure of the smoothing coming from interregional credit, and β_U as a measure of the unsmoothed portion of the shocks.

These interpretations of the β coefficients are extremely important. If they could be sustained, the approach would have the great merit of bringing together in a single framework three smoothing mechanisms which are often treated separately. Sala-i-Martin and Sachs (1992) began econometric work on the attenuation of regional shocks by the central government budget through net transfers, and they have been followed since by a number of authors (see prominently von Hagen 1992 and Bayoumi and Masson 1995). Atkeson and Bayoumi (1993), on their part, considered how much smoothing of regional shocks took place within a country through capital market integration. Quite recently, Bayoumi and Klein (1997) also examined smoothing of regional shocks within a country through borrowing or lending from the rest of the nation. Even more recently, ASY's own work has caught on and inspired at least two major investigations: Athanasoulis and van Wincoop (1998) and Del Negro (1998). Both of these studies, however, center strictly on β_K and β_G, and drop the issue of β_C or smoothing via credit.[1] The narrower focus of all these studies has its own advantages: it permits digging more deeply into some particular aspect. Thus, Athanasoulis and van Wincoop are able to examine smoothing through capital market integration at different horizons, going up to 26 years (in line with their emphasis on *growth* uncertainty). Del Negro, for another, shows how smoothing estimates relating to β_K and β_G vary depending on the statistical measure used to capture permanent income. But ASY's procedure has considerable interest of its own: it allows joint examination of credit, private insurance and public insurance (or public transfers) as smoothing mechanisms. Therefore, the validity of ASY's interpretation of their four β coefficients deserves careful investigation.

There are three reasons why their interpretation may not hold. First, the changes in interregional consumption and saving could stem from movements in intertemporal preferences rather than output. If so, the β_K, β_G, and β_C coefficients would have nothing to do with consumption smoothing. To be more specific, the β_K and β_C coefficients would then concern "crowding in" and "crowding out" rather than smoothing, "crowding in" so far as output was affected by shocks to regional tastes, "crowding out" so far as it was not thus affected. Second, smoothing of regional consumption can take place through accumulation or decumulation of capital within a region and without any interregional borrowing or any income stabilization stemming from interregional property claims. Third, the Miller-Modigliani theorem of the irrelevance of dividend policy might apply: households might see through the corporate veil. Higher corporate saving would then simply induce households to consume more, and β_K and β_C would move in opposite directions. At the limit, the two values could be impossible to estimate separately, and only their sum might be

[1] Compare Lane (1998a).

so. Even if the problem does not go to that extreme, should the two coefficients be highly negatively related on the previous reasoning, the comparative size of β_K and β_C might tell us little about the relative importance of insurance and credit as smoothing mechanisms. All three of the previous problems are consistent with excellent estimates of the β_K, β_G and β_C coefficients.

None the less, ASY provide some persuasive support for their interpretation in the case of the US. First, their estimate of automatic stabilization of 13% by the federal government in the US accords with earlier estimates that were gotten with independent methods. Using US data for gross state product, von Hagen (1992) had found 9-10%. The latter had measured net transfers more narrowly than ASY, including in his construct only personal federal taxes and federal payments to individuals. In related work, where we experimented with a variety of measures of net transfers, we found that if we adopted a measure closer to ASY's and included federal indirect taxes and federal grants to states, we got the same 13% estimate as theirs.[2] These last estimates of stabilization (in the tradition of Sala-i-Martin and Sachs 1992) simply regress net federal transfers (or gross state product minus net federal transfers) on gross state product. The fact that ASY obtained the identical estimate based on a more restrictive specification designating the source of shocks supports their hypothesis that all the shocks come from regional output per head. Had their hypothesis about the source of shocks been gravely mistaken, they might well not have gotten the same estimates (especially since they engage in joint estimation of two other equations through general least squares). We shall encounter other country cases where ASY's specification does not yield the same estimate of stabilization by the central government budget as the one which is found with the more general specification of Sala-i-Martin-Sachs and von Hagen.

In addition, ASY point out that credit should be far more readily available to finance transitory shocks than durable ones, whereas insurance should be able to protect against durable shocks as well as short-lived ones. Hence, if β_K really reflects insurance whereas β_C reflects credit instead, more persistent shocks should lead to higher estimates of β_K relative to β_C. ASY present three sorts of evidence that this is actually the case. First, they show

[2] See Mélitz and Zumer (1998). As we showed in this work, the much higher estimates of 30 to 40% that Sala-i-Martin and Sachs (1992) and Bayoumi and Masson (1995) obtain depend entirely on their use of data for state personal income rather than data for gross state product together with their adoption of the broad measure of net federal transfers in the text. We get the same higher estimates as theirs if we follow their accounting.

that if they average the data over longer intervals, and thus the shocks they measure obtain over longer durations too, the estimates of β_K rise relative to those of β_C. Next, they construct Campbell-Mankiw (1987) measures of persistence of the shocks. When they distinguish between the states with high persistence and the rest, they get higher ratios of β_K/β_C for the sample with the higher measures. Third, they break up the states according to the dominance of agriculture, mineral extraction (mainly oil), and manufacturing in industry; and then they separate each of these three groups between "high" and "low" based on the extent of domination by agriculture, mineral extraction, and manufacturing, as the case may be. Shocks to mineral industries tend to be persistent and property in these industries to be widely dispersed nationally. Thus, they reason that β_K should be larger relative to β_C for the "high" group than the "low" group in the mineral-extraction classification. In the example of the agricultural states, however, the reverse is true: the shocks are mainly short term and the property mostly held locally. Therefore, β_C/β_K should be larger for the "high" than the "low" group in this next classification. In both instances, the expected patterns of β_C/β_K values are confirmed.

On the whole, therefore, ASY offer impressive support for their interpretation of the β coefficients. Nevertheless, we have some objections to their tests and will propose a reformulation of their approach. But before doing so, we should note that we will follow ASY in using the term "risk sharing" to cover interregional smoothing via credit (β_C) as well as insurance (β_K and β_G). When one region borrows from the rest of the nation in order to smooth consumption, there is no risk sharing, properly speaking (unless we insist on the possibility of default). Consequently, the term "risk sharing" can be, and sometimes is, reserved for insurance, or β_K and β_G at present. But as long as regions borrow from others in order to smooth their consumption, the smoothing is still not autonomous, and it does little harm but avoids repetition of cumbersome phrases to follow ASY in adopting the broader, if looser, use of the term "risk sharing" to cover all interregional risk smoothing via β_C as well as β_K and β_G. (See, *inter alia*, Mace 1991, Cochrane 1991, Deaton 1992, and Obstfeld and Rogoff 1996.)

(b) Revised Formulation

ASY give the β coefficients a very precise meaning. Yet they insist on the purely accounting nature of their approach and present their position as based on a mere decomposition of variance. We differ with them on this vital point. As mentioned previously, if the β_K, β_G, and β_C coefficients are to be properly viewed as concerning consumption smoothing, then there must be no shocks to regional tastes. Any taste shocks would either move regional output or regional saving, and neither possibility would brook the interpretation of β_K, β_G, and β_C as relating to smoothing. Thus, the only source of changes in consumption that are consistent with their construction of the β's are output shocks.

In our view, this means that the last member of equations (4) has no place. The only estimate of this equation that would agree with ASY's interpretation of β_K, β_G and β_C would be a perfect fit. Suppose, for example, that the R^2 of the last equation is only 50 percent, and thus half of the variance of $\Delta \log C_i$ is not explained by $\Delta \log Y_i$. Consequently, either this unexplained half of the variance affects the variance of $\Delta \log Y_i$, or it merely reduces the covariance of $\Delta \log Y_i$ with $\Delta \log DI_i - \Delta \log C_i$ and thereby reduces the estimate of β_C, with obvious possible repercussions on β_K and β_G via the identity $\beta_K+\beta_G+\beta_C+\beta_U = 1$. In either event, it is no longer possible to interpret β_K, β_G, and β_C as pertaining strictly to smoothing of output shocks. Moreover, there is every chance that at least some of the unexplained consumption in the estimate of this last equation would come from taste shocks, in which case the idea of any mere smoothing – of output shocks or anything else – is out. Accordingly, we shall view the model as consisting only of the first three members of equations (4), consider $\Delta \log Y_i$ as exogenous, and regard the disturbance terms in these three equations as concerning the composition of smoothing between the three. In view of identity (3), however, the model will also be estimated subject to the restriction $\beta_K+\beta_G+\beta_C = 1-\beta_U$, where β_U is predetermined and inferred directly from the data (in a way that we shall explain).

Two additional considerations argue in favor of our proposed treatment of β_U as predetermined. First, because of identity (3), only three of the four β coefficients can be properly estimated in any event. Second, as we shall go on to show, regression estimates of β_U based on the fourth member of equations (4) bear predominantly on β_C, and therefore yield questionable estimates of this coefficient.

We deviate from ASY in three other respects. First, we take more seriously than they do the aforementioned problem of distinguishing risk *sharing* from autonomous smoothing within a region. Both β_K and β_C could refer to regional accumulation and decumulation without any sharing of risk between regions.[3] In addition, households may modify their consumption plans in the light of business saving, and if they do, β_K and β_C may tell us little about smoothing via insurance as opposed to credit.

[3] Sørensen and Yosha clearly recognize this point in their subsequent paper, Sørensen and Yosha (1998).

Next, we do not follow ASY in treating the uneven sizes of different US states as an issue of heteroskedasticity. Rather than view these uneven sizes as a statistical problem, we see them as an opportunity. The fact that some US states, such as Delaware, are smaller than Ireland, and others, like California and Texas, are easily as big as Spain, makes the US data more pertinent for Europe. The smaller states should be more open, and openness is an interesting variable in itself. The possibility that openness would modify the sources of smoothing seems a fitting subject of investigation in drawing lessons for EMU.[4]

Our final deviation from ASY is the least significant and concerns the treatment of common shocks. As mentioned before, ASY eliminate these shocks from view by introducing a separate dummy variable per year. Sala-i-Martin and Sachs had proposed a different and more economical way to treat the problem, which consists of converting all of the variables into percentages of the national values. In going over this same terrain before (Mélitz and Zumer 1998), we had found the Sala-i-Martin-Sachs method to be efficient. Their technique is just as effective in removing common influences from the analysis as the use of time dummies, but it drastically reduces the number of coefficients that need to be estimated separately. As a result, there are more degrees of freedom and more empirical considerations can be brought into the statistical analysis.

In the light of these differences, the system we propose to estimate will be:

$$\begin{aligned}
\Delta \log y_i - \Delta \log pi_i &= \alpha_K + \beta_K \Delta \log y_i + \gamma_{K,j} (\log X_{i,j}) \Delta \log y_i + \mu_{iK} \\
\Delta \log pi_i - \Delta \log di_i &= \alpha_G + \beta_G \Delta \log y_i + \gamma_{G,j} (\log X_{i,j}) \Delta \log y_i + \mu_{iG} \qquad (5) \\
\Delta \log di_i - \Delta \log c_i &= \alpha_C + \beta_C \Delta \log y_i + \gamma_{C,j} (\log X_{i,j}) \Delta \log y_i + \mu_{iC}
\end{aligned}$$

subject to $\beta_K + \beta_G + \beta_C = 1 - \beta_U, \quad 0 < \beta_U < 1$

and for all j, $j = 1, \ldots, n$, $\gamma_{K,j} + \gamma_{G,j} + \gamma_{C,j} = 0$

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[4] We are also unconvinced by ASY's claim that gross state product "is particularly likely to be measured with error for small states" (p. 1089). There is little evidence that the data for Idaho, Delaware and Rhode Island, for example, is any worse than the one for the large states. To our knowledge, the lower quality of the US regional data for some states – Alaska, Hawai, and the District of Columbia, in particular – can be explained independently of size. "From preliminary estimations", ASY also say, "we found that taking this heteroskedasticity into account had a large impact on the results" (p.1090). But we never encountered the same "large impact".

where the X_j variables are new influences that we will admit into the econometric analysis. The use of lower-case letters instead of upper-case ones in equations (5) is our sign that the variables are now ratios of per capita values in relation to per capita national averages (adding up to one with appropriate weights). The equation for β_U has been replaced by the restriction $\beta_K + \beta_G + \beta_C = 1 - \beta_U$. We have also imposed $\gamma_{K,j} + \gamma_{G,j} + \gamma_{C,j} = 0$ in order to assure that the introduction of the X_j variables does not violate the condition $\beta_K + \beta_G + \beta_C = 1 - \beta_U$. According to our chosen specification, therefore, the sum of the coefficients of $\Delta \log y_i$ across the three equations (5) still yields $\beta_K + \beta_G + \beta_C$, or $1 - \beta_U$, just as before, and the $\gamma_{C,j}(\log X_{i,j})$ terms do not enter. Thus, the X_j variables affect the decomposition of the smoothing without touching the total.

We shall retain four X_j variables in the study. The first relates to the possible smoothing of shocks through capital accumulation or decumulation within a region, and without any recourse to borrowing (+ or –) from the other regions or any reliance on changes in property claims (+ or –) against the others. In order to take the factor into account, we shall introduce a series concerning the regional business cycle as such. The new series, z_i, obtains by first dividing Y_i (the regional level as such) by its own average over the entire period, and then either removing a fitted trend in the series or else using a Hodrick-Prescott filter to get rid of any long run tendency. Since z_i ignores national data entirely, while y_i hinges strictly on regional activity in relation to others, the joint presence of y_i and z_i means that we can interpret any separate significance of z_i as relating strictly to autonomous regional behavior.

A second X_j variable will be regional size, as measured by the ratio of the regional population to the national one, n_i. This variable harks back to our decision to dismiss heteroskedasticity and admit openness. Our specific hypothesis about size, which is associated with the literature on optimum currency areas, is that smaller regions depend more on trade and tend be more open. Accordingly, smaller regions should be more specialized in production, and therefore hold a larger proportion of their property as claims on other regions. On this ground, we expect a negative coefficient γ_K and, consequently, a positive coefficient γ_C.

The Campbell-Mankiw index of persistence, P_i, comes next. Based on ASY's discussion, we anticipate a positive sign of γ_K and a negative sign of γ_C for P_i. In measuring this index, we follow ASY exactly and use three lags.[5]

Fourth and last, we introduce the real interest rate, r (identical across regions). In their comments on different sub-periods of their 1964-1990 sample, ASY suggest that tight monetary policy might be the answer to certain swings in β_C relative to β_K. We shall investigate this hunch directly. While a rise in r associated with tight monetary policy should clearly raise β_K/β_C, a similar rise associated with an increase in factor productivity need not do so. Yet rises in r linked to productivity should not lower β_K relative to β_C either. Therefore, if increases in r are largely associated with tight monetary policy, the general hypothesis of a positive effect of r on β_K/β_C seems reasonable.

II. The US Results

(a) Preliminaries

In doing the work on the US, we stuck to the same sample period as ASY and used the same series as theirs, borrowing any data they constructed directly from them, so as to assure maximum comparability. This meant using their series for consumption, since there exist no consumption data by state. ASY inferred those values from retail sales, as others have done before them.

A few observations about the accounting are in order. The model requires that Y_i minus PI_i relate to business or else only to differences between the location of business and the residence of owners, PI_i minus DI_i relate strictly to central government, and DI_i minus C_i relate strictly to individuals. Foreigners are ignored. State and municipal governments are lumped together with individuals. That is, state and municipal governments' income is included in DI_i and their spending in C_i. As a result, DI_i must include federal grants to states. The most delicate part of the accounting is the need to assure that the difference between gross state product, Y_i, and state personal income, PI_i, has

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[5] Specifically, in constructing the Campbell-Mankiw measure of persistence (P_i), we begin with the AR(3) process, $\Delta \log y_{it} = \mu_i + \sum_{j=1}^{3} \phi_{ij} \Delta \log y_{it-j} + \varepsilon_{it}$, and consequently define P_i for region i as $P_i = \left(1 - \sum_{j=1}^{3} \phi_{ij}\right)^{-1}$.

nothing to do with central-government transfers. Official data for gross state product add up to national GDP, and therefore include corporate income taxes to the federal government and federal excise taxes. On the other hand, official data for personal income by state exclude both taxes. Therefore, any mere subtraction of official data for state personal income from official data for gross state product would necessarily include both taxes, so that if the taxes are to be omitted from Y_i minus PI_i they must be added to state personal income. Yet there are no official decompositions of federal taxes on corporate income or federal excise taxes by state. Consequently, ASY needed to decompose both taxes themselves, and we merely accepted what they did. Correspondingly, the distinction between PI_i and DI_i in their research and ours takes the broadest possible view of the net federal transfers to states.

In order to perform our tests, we required an econometric program of panel data estimation applying to a system of simultaneous equations with correction for covariances across equations and with cross-equation restrictions. Disposing of no appropriate ready-made program, we wrote one ourselves with Stéphane Mysona's considerable help. Table 1 addresses the question of the extent to which the changes in our estimates of ASY's model then result from our modifications in estimation method and our differences in definitions of variables, rather than anything else.

In the first column of the table, we show ASY's estimates of β_K, $\beta_{G,}$ β_C and β_U. Column 2 replicates their estimates using our econometric program rather than theirs, together with their definitions of the variables and their use of separate dummy variables per time period in order to isolate common shocks.

Apart from estimation method, therefore, the only difference of note between the second column and the first one stems from our failure to correct for heteroskedasticity. Since β_K, $\beta_{G,}$ β_C and β_U must sum to one, and we imitated ASY in column 2 by taking none of the coefficients as predetermined, there is a singularity in the system, and our program would not converge. It came to a halt before providing a standard error for β_U and a separate $\overline{R}^2$ for the last of our four equations in system (4). Still, estimates resulted; and as can be seen, the differences between our estimates of β_K, $\beta_{G,}$ β_C and β_U in column 2 and theirs in column 1 are negligible.

The last two columns concern the effect of converting our variables into ratios and therefore dropping the time dummies. In both columns, we show simple pooling estimates. We carried out "within" estimates as well, but the two yield nearly identical results. In the first of these next two columns (column 3),

β_U is estimated as before in column 2, whereas in the last one (4), β_U is predetermined, and $\beta_K + \beta_G + \beta_C$ is constrained to equal $1 - \beta_U$. Once again, in column 3, where we estimate all four β's and fail to recognize the presence of a singularity, the program would not converge, and incomplete results follow. The last column shows the results when we introduce our predetermined value of β_U. Our choice of this value obviously requires separate discussion at this point.

We constructed β_U there, as we shall throughout, by calculating the variance of consumption C_i (not c_i) and dividing by the variance of output Y_i (not y_i) every year and then averaging the individual (β_{Ut}) values over all the years. Our main reason for doing so is our finding, in earlier experiments, that this measure offered us the lowest value of β_U (the closest to ASY's to boot), and therefore the widest scope for the application of ASY's method of decomposition. Among other measures of β_U, we tried logs, first differences, ratios instead of levels (c_i and y_i instead of C_i and Y_i), variances of regional time series (β_{Ui}) rather than variances of regional cross-sections (β_{Ut}). But all the other measures yielded higher values of β_U. Very significantly too, our choice makes sense: if movements in regional consumption stem exclusively from movements in regional output, as the model says, then any lower cross-sectional variance of regional consumption than cross-sectional variance of regional output must reflect smoothing.

Our measure of β_U is 39% for the US in 1963-1990. This is an interesting statistic in itself, lending support to the notion that much cross-regional consumption smoothing takes place within the country. Yet the ratio is higher than the estimated β_U in the preceding columns of the table.

In light of the results in columns 3 and 4, it is clear that our conversion of the variables, as such, does not account for significant differences between our estimates and those of ASY. But we must correct any possible impression, based on Table 1, that even our decision to treat β_U as predetermined makes little difference. The calculated β_U (on any of our aforementioned measures) and the estimated one can be far apart, and if they are, the difference between the two β_U values will be mostly compensated by an opposite change in β_C. A glint of this may already be gotten from Table 1 by taking a close look at the difference between columns 3 and 4. But Table 2 makes the point transparent.

There we show ASY's aforementioned results when they estimate the β's based on first-differences over successively longer observation periods. The first column contains their estimates based on first-differences in annual observations, and the next three those based on first-differences with successive 3-year, 5-year, and 10-year intervals between observations. As can be seen, with the lengthening of the time period, β_U rises systematically while β_C moves correspondingly downward, even becoming negative. Column 5 carries the process that ASY began in columns 1 through 4 to its logical conclusion by providing the "between" estimates of the coefficients, which are simply the cross-sectional estimates based on the averages (of the annual first-differences) over the entire sample period. In these next estimates, however, we use our definitions of the variables and our test procedure rather than ASY's, since we know now that doing so makes little difference. As we see, the estimate of β_U goes up all the way to 0.79 in column 5 and, correspondingly, β_C becomes even more negative than before. But if we calculate β_U from the data itself in our previous manner, we get an observed β_U (based on the averages for C_i and Y_i over 1963-1990 as a whole) of only 0.31, or less than 0.39, the previous value in Table 1, and a far cry from 0.79. The last column in Table 2 presents our revised "between" estimate resulting from the constraint $\beta_K+\beta_G+\beta_C = 1-\beta_U$ with β_U equal 0.31. In this case, β_C rises substantially relative to the previous column, and in line with the earlier estimates in the first four columns. It is thus clear that regression estimates of β_U bear mostly on β_C and render the estimates of β_C of little interest. Very significantly, we found this generally true, in dealing with other national data sets as well: regression estimates of β_U will sometimes deviate markedly from measured values and when they do, β_C takes the brunt.

But this is not how ASY interpret the matter. They take their figures for β_C in the first four columns of Table 2 seriously, and refer to possible "dis-smoothing" of consumption as "lenders actually pull out loans from states that have been unlucky for several years in a row" (p.1097). Accordingly, they offer the results of these columns as supporting their major conclusion that as shocks become more durable, regions rely more heavily on insurance relative to credit. Evidently we disagree. Yet ASY may still be right on the basic point about the effect of durable shocks in raising β_K relative to β_C. Though β_U in the last column, or 6, is down to 0.31, the ratio β_K/β_C is still much higher there than in column 1. We will return to the issue.

(b) The Main Tests

With these preliminaries aside, we may turn to the results of testing our equation system (5). Table 3 presents these results when all four X_j variables – z_i, n_i, P_i, and r – are included. We experimented with several measures of z_i and r, and therefore those serving in the table should be mentioned at once. For z_i, we report on the coefficients resting on the Hodrick-Prescott filter.

However, those based on the elimination of a trend are almost identical. The measure of r in the table is the short term interest rate in the OECD *Economic Outlook* (corrected for CPI inflation). But it makes little difference if we use a bank-loan rate or a one-year security rate instead.

Let us focus first on the revised estimates of β_K , β_G and β_C. As regards β_G, or smoothing via central government net transfers, there is no difference to speak of. The new estimate of 13% is identical to ASY's. But the revised estimates for β_K and β_C differ widely from theirs. β_K is now equal to β_C. Thus, smoothing via credit becomes as high as smoothing via insurance. This change is entirely attributable to z_i, or the admission of autonomous smoothing. Removing z_i from the equation brings us back to estimates of β_K and β_C nearly identical to those in Table 1, but leaves the rest of Table 3 essentially unchanged. Conformably, we see from $\gamma_K(z)$, $\gamma_G(z)$, and $\gamma_C(z)$ that z_i reduces insurance a lot (by 0.07) and raises credit commensurably while leaving net transfers unaffected. It is apparent, therefore, that failure to take self-financing into account in analyzing consump-tion smoothing leads to exaggerated estimates of interregional insurance (via movements in interregional claims on income and property) and underestimates of interregional borrowing. This conclusion will be corroborated again in dealing with Canada.[6]

As regards n, we do not get confirmation of our hypothesis that largeness, or lower openness, reduces interregional portfolio diversification per head. But an unanticipated result emerges in Table 3. Based on $\gamma_G(n)$, it seems that larger US states receive (pay) larger net transfers from (to) the central government in response to adverse (favorable) regional shocks. These net transfer payments appear to be at the expense of credit rather than insurance, as they are matched by a negative value of $\gamma_C(n)$, not $\gamma_K(n)$.

The effect of the Campbell-Mankiw persistence index supports ASY entirely. Higher persistence raises reliance on risk sharing via insurance relative to risk sharing via borrowing. Indeed, the Student t's associated with P_i are the highest in Table 3. This is stronger corroborative evidence of ASY's interpretation than the one of the "between" estimate that we noted earlier in connection with the last column of Table 2, since those earlier results can now be seen to be partly flawed by a certain confusion of regional and interregional smoothing. The problem cannot be resolved in the "between" estimates where

[6] The conclusion is relevant in judging the results of Athanasoulis and van Wincoop (1998) and Del Negro (1998) as well, both of whom also fail to control for intraregional adjustments, and as a result, we think, also exaggerate the importance of interregional insurance.

z_i is impossible to introduce. Still, β_K rises too much relative to β_C in column 6 of Table 2, as compared with column 1 of the same table, for this earlier evidence in favor of ASY's hypothesis to be totally dismissed.[7]

Finally, the estimates of the impact of the real interest rate on smoothing activity confirm ASY's hypothesis that higher real interest rates reduce reliance on interregional credit in favor of asset diversification.

Table 4 reports one further "between" estimate for the US: namely, the one resulting by adding the size of states to the earlier "between" estimate of Table 2 with a constrained β_U of 0.31 (which is repeated on the left side of the table for convenience). The indices of cyclical behavior and persistence of shocks are left out since these variables have no place in a "between" estimate (where there are no time series).[8] The interest rate is omitted as well, since we treated this variable as identical across regions. Quite notably, Table 4 shows $\gamma_K(n)$ as immensely significant with the expected negative sign. The estimate of $\gamma_G(n)$ in the table also indicates that central-government transfers favor small rather than large states in the event of an adverse durable shock (just the opposite of what we found with respect to transitory shocks in Table 3). But the result regarding $\gamma_K(n)$ bears most emphasis. It confirms the idea that openness promotes the cross-regional ownership of property and the associated insurance. Upon reflection, our inability to obtain this result earlier and our ability only to find it now may seem natural. Openness is an institutional feature, whose effects might only be possible to detect in cross-sectional estimates focusing on behavior over a long period of time. Admittedly, the estimates of $\gamma_K(n)$ and $\gamma_C(n)$ in Table 4 continue to confuse regional and interregional behavior. But it is difficult to see why the smallness of regions would promote smoothing via insurance relative to credit except for interregional effects.

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[7] We have acquired doubts, however, about ASY's evidence based on the industrial decomposition of agriculture, manufacturing, and mineral extraction. Those results do not hold up in our tests for "low" relative to "high" agricultural states, though they do so for "low" relative to "high" mineral-extraction states. Upon reflection, we also question the comparison of the smoothing by different states in the same industrial classification. Why not compare all of the agricultural states with all of the mineral-extraction states, or for that matter, with all of the rest? However, the importance of the point must not be exaggerated, since the evidence based on the persistence index largely covers the same ground.

[8] By their very nature, however, these estimates focus on persistent rather than temporary shocks, and thereby largely take persistence into account.

(c) Interregional relative to regional market smoothing

Before leaving the US, we must come back to the question of the total interregional smoothing through market forces and through the centralized budget. The matter has now become more complicated. We saw earlier that β_K and β_C, which reflect smoothing through market behavior, contain both regional and interregional smoothing in ASY's original approach (Table 1). Now that we have separately investigated regional smoothing, where does the matter stand? In fact, it is still true in Table 3, as would be the case with the same figures for the β's in Table 1, that of every dollar shock to regional output (in relation to the rest of the nation), 39 cents is not smoothed, 13 cents is smoothed interregionally through net government transfers, and the other 48 cents of market smoothing may be achieved either regionally or interrregionally. What we have done is simply to correct for strictly regional smoothing in dividing up the same 48 cents between β_K and β_C. As a result, the ratio of β_K to β_C in Table 3 now gives a better measure of the relative importance of insurance and borrowing *in interregional smoothing* than the earlier ratio of the two in Table 1. However, the proper division of the 48 cents between *interregional and regional* smoothing remains an open question.

Still, the estimates of $\gamma_K(z)$ and $\gamma_C(z)$ shed some light on the extent of interregional smoothing as such. According to these estimates, in the absence of any correction for regional smoothing, β_K would be approximately 12% (–0.07 times -1.67) higher (see the next footnote) and β_C 12% lower. In other words, based on Table 3, β_K would be approximately 36% and β_C 12% rather than both 24%. Take the extreme case where all of the strictly regional smoothing is reflected in business saving. In that case, 12/24 or 1/2 of the 48 cents of market smoothing would need to be done regionally to explain the rise of β_K of 12% and the accompanying equivalent fall of β_C. That would then leave total interregional smoothing by market forces of 24 cents (instead of 48) in relation to 13 cents by the government. But this 24 cents is an upper limit. In fact, interregional smoothing could account for as little as 13 cents, or the level attributable to upper-level government transfers. But to do so, the interregional smoothing would need to represent only 13/48 of the total smoothing by market forces, or about one quarter of the total.[9] Thus, in light of our

[9] To explain our calculations, let the fraction of total smoothing of the 48 cents, or by market forces, that is done *interregionally* be x. Then we have

$$x\,\beta_K + (1\text{-}x)\,a = \beta_K + \gamma_K(z)\,\overline{\log z_i}$$

$$x\,\beta_C + (1\text{-}x)\,b = \beta_C + \gamma_C(z)\,\overline{\log z_i}$$

estimates, the predominance of market forces in interregional smoothing remains probable. But it is lower than ASY maintained, and we simply cannot pin an exact number on the division of the risk *sharing* between the market forces and government transfer payments.

III. The Canadian, British and Italian Evidence

Regional consumption data does not exist for the US and the numbers needed to be inferred from retail sales. This makes it important to experiment with the ASY approach in countries where regional consumption data does exist. Canada, the UK, and Italy are three such countries. The essential required data to apply the ASY approach is also available for these three countries. What are the results?

Table 5 begins the discussion of these other countries with some preliminary statistics. As we see from the number of regions and the length of the sample periods, the total number of observations for Canada, the UK and Italy is only roughly 200 to 300, as opposed to 1300 for the US. This in itself could lead to a lower quality of estimates. The table also shows considerable differences in the calculated β_U values among the three countries. They go from 55 to 60% for Italy and the UK to only 37% for Canada. Even the 55-60% Italian-British level implies substantial smoothing, though how much of it takes place through interregional activity is obviously an open question. But the low Canadian figure, roughly matching the one for the US, virtually assures us substantial interregional smoothing in Canada.

Those sizable differences for β_U between North America and Europe are worthy of contemplation of themselves. Can it really be that there is much more interregional smoothing in Canada and the US than in the UK and Italy? Or is it instead that in the federally organized and geographically sprawling two North American countries, the decomposition into regions makes more sense

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where a and b regard the total market smoothing via business saving and household saving, respectively, which is strictly *regional*, and $\overline{\log z_i}$ is the mean of log z_i in the data sample. Since $\beta_K = 0.24$, $\gamma_K(z) = -0.07$, $\beta_C = 0.24$, and $\gamma_C(z) = 0.07$ in Table 6, and $\overline{\log z_i} = -1.67$, if all of the smoothing that took place regionally were done via business saving and b were zero, x would be approximately 1/2 (as mentioned in the text) and a would be 0.48. For x to be 13/48, which would render interregional smoothing through market forces no higher than β_G, b must be $\cong$ 1/13 and a $\cong$ 2/5. Note that the estimates of $\gamma_K(z)$ and $\gamma_C(z)$ guarantee that the lion's share of the regional smoothing (that is, the lion's share of the fraction 1–x) is done by business saving, and therefore a is much larger than b.

than in the politically unified and smaller European ones? One factor that would plead in favor of the latter interpretation would be evidence that the movement in the regional composition of output is much smaller in the European cases. The last column of Table 5 addresses this question. It shows the coefficients of variance for regional output per capita (interregional variance divided by national average) for all four countries. Based on this column, there is indeed evidence of lower variance of regional output for the UK than North America but not for Italy.

Let us begin examining the test results with the simple ASY framework given our definitions and our estimation procedure ($\beta_K + \beta_G + \beta_C = 1 - \beta_U$). As we see from Table 6, only for Canada are β_K, β_G, and β_C significant (and the same is true regardless of "within" or pooling estimates). The British and Italian results resemble one another. In either case, β_G and β_C are totally insignificant in the pooling estimates, and all three coefficients only matter statistically in the "between" ones. The "between" estimates of β_G for the UK and Italy of 9 and 13%, respectively, are acceptable, if on the low side. On the other hand, the corresponding β_G estimate for Canada of zero conflicts with independent knowledge and other estimates.[10]

Efforts to improve the previous estimates by introducing conditioning influences on the β coefficients, in our previous manner, proved successful only for Canada. Table 7 reports the effect of admitting the regional business cycle, z_i, in all three countries. As can be seen, Canada is the only one where the $\gamma(z)$ coefficients are significant. Moreover, introducing this regional variable has the familiar result of reducing β_K in favor of β_C. Indeed, in the Canadian case, once we take into consideration the regional business cycle as such, β_C even exceeds β_K. Something similar occurs in the UK, though not in Italy (where the brevity of the observation period renders the z_i variable questionable). But perhaps little should be made of this last British result since the $\gamma(z)$ coefficients in this country are insignificant. As regards both the UK

[10] Based on different accounting, Bayoumi and Masson (1995) showed 18% (just as we did based on their accounting (Mélitz and Zumer 1998)). When we used the identical accounting as the one in the text, we got 23%. These wide differences in estimates show that the ASY specification will not always give the same results for the role of net transfers from the central government as those coming from the less restrictive specification in the tradition of Sala-i-Martin and Sachs (1992). It should be noted too that the β_G estimates for Italy and the UK are underestimates as compared to the US and Canadian ones, because they are based on narrower definitions of net transfers from the central government. See the Data Appendix.

and Italy, we can only conclude that β_K and β_C tell us little or nothing about smoothing via insurance or via credit.

In Table 8, we focus exclusively on Canada. There we show what happens when all of the X_j variables are added to the analysis. The measures of persistence and the real interest rate are the same as in the earlier tables for the US. Persistence has the right positive effect on β_K and negative effect on β_C, which we regard as support for the interpretation of β_K and β_C as concerning insurance and credit, respectively. On the other hand, the effect of the real interest rate on credit relative to insurance does not show up. The only other notable result in Table 8 is the negative effect of region size on β_G, which is matched by a positive effect on β_C. There is therefore some sign that smaller provinces get more help from the federal government in the event of adverse shocks, and correspondingly borrow less.

While broadly adequate, the quality of the Canadian results is still below that of the US estimates. One of our efforts to find out why seems to us of interest, though the exercise proved unsuccessful. Courchene and Laberge (1998) show that the US is as important to most individual Canadian provinces as a trading partner than the rest of Canada. Could it be therefore that Canada can only be properly viewed as an integrated economy when the US is included? Based on this query, we added the US as a region in the analysis, assigning a weight of 40% to it in "Canada-US" as a whole, just enough to make the US as important to Ontario and Quebec as the rest of Canada. In order to do so, we converted the US per capita figures into Canadian dollars by using the nominal Canadian/US dollar as adjusted for relative CPI inflation between Canada and the US. Following, we recalculated all of the per capita "Canadian" regional figures accordingly. (We also experimented with the real exchange rate as a separate X_j variable – admitting the constraint $\gamma_K + \gamma_G + \gamma_C = 0$ – because of possible associated distortions.) The results were always worse than before.

The failure of the model for the UK and Italy remains a disappointment. In the case of the UK, the only reasons we can see are those we mentioned before: too few observations and too small an amplitude of regional shocks. As concerns the Italian flop, the brevity of the sample is our only explanation.

IV. The International Dimension

The possible international application of the ASY approach is important if we wish to draw lessons for EMU, since this new monetary order will mean moving from international monetary relations with other members of EMU to essentially domestic monetary relations with them. In proceeding with the effort, it should be borne in mind that national statistics are superior to regional ones in many ways. Current account balances offer figures for net foreign borrowing as such. Similarly, the differences between gross national product and gross domestic product record the actual flows of net factor income from abroad. In some respects, therefore, the international application of the model should be easier than the national one. Risk *sharing* will be simpler to separate from strictly domestic responses to idiosyncratic risk, while some of the exact sources of the *sharing* will be identifiable directly. In addition, we will not need to worry about net transfers through a supra-national government agency, since no such agency exists, except possibly in the case of the European Union, where its significance is small and can be gauged independently.

One major gap nevertheless prevails in national statistics: these do not measure gains and losses on net foreign assets. The differences between GDP and GNP cover recorded income flows but leave capital gains and losses out of account. As Obstfeld (1986, pp. 82-86) and Stockman and Svensson (1987) have separately emphasized, this omission may matter greatly in analyzing market behavior. Consequently, we will need to continue relying on the earlier kind of inferences.[11] The home-country bias in international portfolios may diminish the problem. But long-term gross capital movements are considerable, and unrecorded capital gains and losses on foreign positions can be huge.

The accounting identity we propose to use is the following:

$$Y_i = \frac{Y_i}{GNP_i} \frac{GNP_i}{A_i} \frac{A_i}{C_i} C_i \qquad (6)$$

where Y_i is gross domestic product (as before), GNP_i is gross national product, A_i is home absorption, so that $Y_i - A_i$ is exactly the export surplus on current account, and C_i is the sum of private and public consumption. In accordance with identity (6), we then propose the following adapted version of the ASY model:

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[11] The same difficulty was present in the national applications, but it was less important there since we needed to rely on the earlier sorts of inferences independently.

$$\Delta \log y_i - \Delta \log gnp_i = \alpha_{K1} + \beta_{K1} \Delta \log y_i + \gamma_{K1,j} (\log X_{i,j}) \Delta \log y_i + \mu_{iK1}$$
$$\Delta \log gnp_i - \Delta \log a_i = \alpha_C + \beta_C \Delta \log y_i + \gamma_{C,j} (\log X_{i,j}) \Delta \log y_i + \mu_{iC} \qquad (7)$$
$$\Delta \log a_i - \Delta \log c_i = \alpha_{K2} + \beta_{K2} \Delta \log y_i + \gamma_{K2,j} (\log X_{i,j}) \Delta \log y_i + \mu_{iK2}$$

subject to $\beta_{K1} + \beta_C + \beta_{K2} = 1 - \beta_U, \quad 0 < \beta_U < 1$

and $\gamma_{K1,j} + \gamma_{C,j} + \gamma_{K2,j} = 0$

Once again, small letters designate ratios, or at this point per capita national values divided by per capita international ones. We continue using β_C to refer to smoothing via credit and β_K via insurance. Thus, β_C now occurs in the second equation instead of the third, two β_K's (β_{K1} and β_{K2}) are necessary, and there is no β_G. The coefficient β_{K1} regards risk sharing via income flows (first equation), and β_{K2} would encompass any risk sharing via capital gains and losses (third equation). Of course, how much risk *sharing* is actually implied by the third equation rather than risk smoothing at home is not obvious, to say the least. This next equation can best be seen as referring broadly to smoothing of idiosyncratic shocks through domestic saving coming from all possible sources except those in the first two equations, foreign factor income and foreign borrowing. Only success with additional variables X_j could possibly warrant the inference that β_{K2} relates partly to risk sharing through net property claims on foreigners. We shall rely heavily on the domestic business cycle, z_i, in this respect. Any significance of z_i will be interpreted to reflect strictly national behavior, and if modifying the estimate of β_{K2}, as bolstering the idea that β_{K2} signifies partly international risk sharing (but how much?).

In order to study equations (7), we start with the largest possible sample of OECD data. Our data covers 23 countries over 1970-94, or 575 observations. The countries include all 15 members of the EU plus the US, Canada, Japan, Australia, New Zealand, Iceland, Norway and Switzerland. Except for GNP (and for consumer durables in those of our experiments where we remove these goods from C), all of the series are available in real terms and do not need to be deflated. 1990 exchange rates serve to convert all the data into US dollars. In order to arrive at per capita values over the 23 countries, we used either GDP weights or trade weights (reflecting real exports and real imports of goods *and services*). Since both sets of weights give identical results, we retain the trade weights, which appear preferable to us. Size seems especially related to trade volume so far as international risk sharing is concerned. It is imaginable, of course, that in the international application, keeping the national values intact and using time dummies to correct for common influences would be better than dividing by international aggregates.

For this reason, we estimated both ways. But there was no fundamental difference between the two (as had been true for the US). Therefore we shall continue to report results based on the ratios and without time dummies.[12]

In a sense, we were lucky. The model might have been inapplicable. Everything depends on the presence of a β_U less than one or some smoothing: otherwise, there is nothing to say. Fortunately, the cross-sectional variance of national consumption per capita (the actual value, not the ratio) is lower than that of national Y per capita in the OECD sample both for all of the OECD23 and for the EU15 sub-sample. The measures of β_U that rest on pooling of annual observations and averages over all the years (the "between" calculations) are both the same: 80% for the OECD23 and 77% for the EU members. Hence, there is 20% smoothing to be analyzed in one case, 23% in the other. Those numbers for smoothing are lower than the previous ones in the national samples, and therefore, to all evidence, the higher economic integration within countries than between them promotes smoothing (compare Crucini 1998). We also found higher variances of the national y_i's than the regional y_i's in the individual countries: more exactly, higher coefficients of variance for the former than the latter on the average for the different dates. Table 9 provides the exact numbers. These are at least four times higher nationally than regionally. In this table, we also show the drop in the coefficients of variation when real effective exchange rates are kept the same.[13] Those corrected figures may be more comparable to the earlier ones in Table 5, since in dealing with countries, it was not possible to adjust for differences in inflation rates between individual regions and therefore the earlier

[12] In fact, in correcting for the common influences by taking time into account and without conversion into ratios, we used "within" estimates instead of time dummies. That is, we simply deducted the contemporary cross-country means of the national values from the observations (in log form) at all dates before estimating. In this manner, we did away with the need to estimate separate coefficients for time dummies (per equation) at all separate dates. This greatly simplifies the econometric analysis and permits estimates of the γ coefficients just as easily as using ratios. Ratios are nevertheless preferable. Not only do they perform as well in eliminating the common influences, but they permit subsequent use of "within" in the conventional way: to correct for constants in the cross-sectional dimension (per country). For example, the conversion into ratios permits adjusting later on for the higher rate of growth of Greece than the UK during the sample period. As it turns out, this next advantage is unimportant in the study since the pooling and "within" estimates are about the same. But this need not have been the case: using the ratios got us the extra information very cheaply.

[13] Failure to correct for changes in effective exchange rates allows movements in exchange rates to infiltrate the data (as well as differential inflation), despite the conversion into dollars, because of movements in exchange rates between third or non-dollar currencies.

figures were biased downward relative to the international ones.[14] Whether we look at the adjusted or unadjusted international figures, however, the picture is the same: the idiosyncratic shocks are immensely higher between nations than within nations.[15]

Table 10 contains the essential results of the international tests. Our experiments with the elimination of consumer durables from (private and public) consumption did not yield any differences, and therefore are not reported.[16] The Campbell-Mankiw persistence index P and the real interest rate are omitted since these two variables never emerged as significant. While the failure of both is a disappointment, it is not as damaging as it would have been in the national examples, or when the failure occurred in the previous section, for the reasons we mentioned. The business cycle variable, z_i, emerges as important in the pooling equations for the OECD but not the EU ones, and even in the case of the OECD, the presence of z_i does not change the estimates

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14 Del Negro (1998) has succeeded in constructing price indices for individual states of the US recently, but did not find his results to be affected (compare Athanasoulis and van Wincoop 1998). See also Hess and Shin (1998).

15 We may comment at this point on the well-known "quantity anomaly", which exists in the time dimension as opposed to the cross-sectional one. The "anomaly" says that if we compare the correlations over time between per capita consumption in different countries with those between per capita output in the same countries, the former correlations are lower, whereas on frequent theoretical assumptions, they should be higher. (See Backus, Kehoe and Kydland 1992, Obstfeld 1994, and Obstfeld and Rogoff 1996, ch. 5). A few studies have also found the "anomaly" to hold within countries (for a general review, see Crucini and Hess 1999). Not surprisingly, the major anomalous results of the earlier studies are found here too. The consumption correlations [$r(\Delta \log C_i, \Delta \log C_{US})$] in the OECD are lower on average than the output correlations [$r(\Delta \log Y_i, \Delta \log Y_{US})$]: 11% in one case as opposed to 16% in the other (with the US serving as the base country). The same follows within the US: 42% on average as opposed to 67% (with California serving as the base). (The choice of base country/region affects the levels, but not the relative order of the two correlations.) The results for our other samples vary. As regards the EU15 and the UK, we find the two correlations to be equal (78% in relation to Germany; 91% in relation to the "Southeast"). For Canada and Italy, we find the correlations to go the right way, or to be higher for consumption (92% as opposed to 72% with respect to Ontario; 92% as opposed to 71% in relation to Lombardy). But the whole anomaly depends on special assumptions (about the structure of preferences for a narrow variety of goods, excluding leisure) that are stronger than any we require.

16 One difference between the accounting in the international study and the earlier domestic one should be noted. Though investment by lower-level governments was present in consumption for the US and Canada before, all government investment is now excluded from consumption. The earlier inclusion of lower-level government investment stemmed from a desire to isolate the regional smoothing by the upper-level government budget as such.

of the β coefficients much. Hence, it is not clear how much β_{K2} can be associated with international risk sharing.

A few strong conclusions nevertheless surface. Movements in net foreign factor income turn out to be far more important than foreign borrowing in the smoothing of shocks internationally. This is true both in the short run and the long run. The higher significance of these income movements than foreign borrowing is particularly marked for the EU15, where the current account balance is never important. Foreign borrowing is clearly significant only for the 23 OECD countries in the long run (the "between" equation). The usual stress on current account balances in formal macroeconomic analysis of stabilization may therefore be overdone. The comparison between β_{K1} and β_{K2} is also of interest. β_{K2} is more important than β_{K1} in the short run, but the relative importance of β_{K1} climbs in the long run. Within the EU15, β_{K1} even dominates β_{K2} in the long run – an impressive result since β_{K1} has a very narrow interpretation and β_{K2} a very broad one. The general picture in the long run for the EU is interesting: more than half of the smoothing comes through risk *sharing*, and all of this risk sharing concerns insurance (diversified property holding) rather than credit.

Some important results also relate to n_i. The variable refers to openness in Table 10, as measured by the ratio of exports plus imports to GDP, rather than relative population, which had been essentially a proxy for openness before. (Note carefully in this connection that a rise in n_i now means more openness, whereas it meant less openness before.) Once again, as was true for the US, n_i emerges as important with the right signs in the "between" estimates or those relating to the long run. Openness should lead to more cross-ownership of resources, and therefore positive values of $\gamma_{K1}(n)$ and $\gamma_{K2}(n)$ and an offsetting negative one of $\gamma_C(n)$. That is exactly what happens in the estimates.[17] The values of the $\gamma(n)$ coefficients are significantly higher within the EU15 than the OECD23, which would indicate that the impact of openness is more pronounced between the 15 EU members than the 23 OECD ones. The right sign of $\gamma_{K2}(n)$ in the "between" equation also tends to reinforce the interpretation of β_{K2} as pertaining partly to international risk sharing. Nevertheless, at least for the EU15, the Student t for $\gamma_{K2}(n)$ is much lower than the one for either $\gamma_{K1}(n)$ or $\gamma_C(n)$.[18]

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[17] Compare Lane (1998b), who gets corroborative results.

[18] Sørensen and Yosha (1998) provide an international extension of ASY of their own, which diverges widely from ours. In their main analysis, they start from the identity

V. Conclusion

We have investigated the channels of interregional and international smoothing of output shocks in a framework originated by ASY, to which we have added modifications. In opposition to the innovators, we interpret the model as postulating that all regional (national) movements in consumption relative to the national (international) average stem from output shocks, and as otherwise inapplicable. Accordingly, we treat the degree of unsmoothed consumption as predetermined in the model. We also try to enrich the model by incorporating additional influences, among them, the openness of individual regions/countries. The model stands up very well in the US in its revised state. As regards the US, our estimate of regional stabilization through fiscal federalism is identical to ASY's. But we come up with different estimates of risk sharing via the two market mechanisms than theirs. Whereas they had found insurance to exceed credit greatly, according to our results, credit is as significant as insurance in interregional risk sharing. We also find these two market channels of interregional smoothing, taken together, not to dominate the smoothing through the federal-government budget nearly as much as ASY had proposed. These conclusions of ours, which depend on our decision to admit autonomous smoothing by regions, are corroborated by the Canadian results. On the whole, our estimates in the US and Canada are remarkably similar, though the model works notably better for the US. However, the model performs badly in the UK and Italy, for reasons which we do not pretend to master.

The adaptation of the model in the international case bore some fruit as well. The raw data suggests different orders of magnitude for shocks and their smoothing at the international level and the national one. The idiosyncratic shocks are larger and the smoothing is lower internationally. The econometric

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$$Y = \frac{Y}{GNP}\frac{GNP}{NI}\frac{NI}{DI}\frac{DI}{C}C$$

where all the variables have the obvious designations, and therefore make no use of the distinction between Y and A or current account balances in this part. Rather, they introduce separate consideration of the current account balance late in the paper, in a passage where they focus strictly on the distinction between domestic and foreign investment. In addition, they continue, as in their national applications, to distinguish between smoothing behavior by firms (GNP/NI), through net public transfers (NI/DI) (while correctly using OECD definitions, which relate the differences between national income, NI, and disposable income, DI, strictly to *international* taxes and transfers, as Sørensen and Yosha require), and by households (DI/C). Most important, they mostly come up with much lower estimates of β_U than our calculated ones. Their estimates go down to 56% and average roughly 65%. When confronted with evidence that consumption varies more than these estimates signify, Sørensen and Yosha appeal to taste shocks. But as seen from our earlier discussion, we believe that this avenue of reconciliation is not open to them.

analysis reveals further major differences at the national and international levels. Most important, credit plays a much smaller role relative to claims on property (to labor income as well as wealth) in risk sharing between countries, especially in the long run. As another major conclusion, openness matters in the long run. The role of openness emerges clearly not only in the international evidence but the US one as well. Both sets of evidence support the hypothesis that openness promotes risk sharing via insurance as opposed to credit. On this ground, openness can be said to increase protection, since insurance is clearly more susceptible than credit to provide cover against durable shocks. On the other hand, of course, openness might amplify the shocks themselves, as we will mention again.

What are the implications for EMU? Professional debate tends to emphasize the fact that members will sacrifice independent monetary policy. Yet, based on our general approach, about 75-80% of idiosyncratic output shocks go unsmoothed in the EU countries. Therefore the importance of the sacrifice can be exaggerated. Still, it remains true that part of the smoothing showing up in our tests could stem from monetary policy. In particular, the β_{K2} coefficient in our pooling equations could be partly the work of monetary policy, which can affect domestic saving. No smoothing effect of monetary policy via the current account emerges in our tests, since β_C is not significant in our pooling estimates (the essential ones in point, in light of the fact that monetary policy acts mainly in the short run).[19]

Quite significantly, though, our results support two reasons to expect more smoothing through market forces under EMU. First, our estimates of β_C are higher for the US and Canada than for the OECD and the EU, and they are so not only in absolute terms but relative to $1-\beta_U$ (or total smoothing). The natural interpretation would be that regions are able to borrow more easily from the rest of the country than countries can from the rest of the world.[20] If so, credit should become more readily available to finance temporary problems in the EMU than it is now among the member countries of the system. Secondly, our results about openness show that in the long run, economic integration favors the holding of property claims across borders. For this reason, EMU might be expected to progressively lead to more insurance against shocks. In

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[19] However, it could also be argued that, in any event, our tests are not well designed to display the stabilizing effect of monetary policy via the current account, because this stabilizing effect works by promoting an export surplus or foreign lending during a recession, whereas on our construction, this last line of influence would be interpreted as destabilizing, since it means lowering current consumption at the time of an adverse shock.

[20] This is clearly another manifestation of the Feldstein-Horioka (1980) result.

conformity with both of these reasons for additional smoothing under EMU, there is indeed greater smoothing of idiosyncratic shocks through market forces within countries than between countries.[21]

Another frequent criticism of EMU concerns the absence – or near-absence – of a mechanism of net public transfers such as the one which exists within countries and that we have confirmed for the US and Canada (perhaps even for Italy and the UK, if we base ourselves on the "between" estimates for these two countries). But this argument against EMU depends on the principle that some aggregate smoothing capacity will be lost under EMU and will need to be replaced. Given the previous reasoning, however, this principle is not necessarily correct: while monetary union will possibly reduce smoothing by eliminating monetary independence, it will increase smoothing through market channels.

Another major dimension of the broad issue under discussion, to which we have already alluded, is the impact of EMU on the importance of the idiosyncratic shocks themselves rather than the percentage of those shocks that are smoothed. On this matter, there are arguments going both ways. Some studies consider that trade integration within EMU will reduce the asymmetric components of business cycles in the EU. Krugman (1993) has famously argued that monetary unification will promote regional specialization and thereby increase the idiosyncratic element in the shocks the country members face. Frankel and Rose (1997) come to the opposite conclusion based on the evidence.[22] But the issue falls outside the ken of our investigation. We take the amplitude of shocks for granted and can only speak about the extent to which they are smoothed and how. Within those limits, the fundamental lesson is the earlier one: based on our extension of the imaginative work of Asdrubali, Sørensen and Yosha, EMU will encourage smoothing through market channels.

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[21] Hess and Shin (1997) come to a similar conclusion about the implications of EMU in a related summary of the evidence.

[22] Research is active on the question. See Artis and Zhang (1997) and Imbs (1998).

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Data Appendix

United States

We have deliberately used the same annual data that was employed by Asdrubali, Sørensen and Yosha (1996), much of which we obtained directly from them. The data covers their study period, 1963-1990, and includes all 50 states but not Washington DC. With respect to the only variables present in our study of the US that they did not use, the nominal interest series comes from the OECD *Economic Outlook* and the consumer price index (serving to correct the interest rate for inflation) comes from the OECD *National Accounts Tables*. These last data sources were employed as well in constructing the short term real interest rate for Canada, Italy and the United Kingdom

Canada

For Canada, all of the main series come from an extract of the CANSIM database "Provincial Economic Accounts”, obtained directly from Statistics Canada, and they cover 1961-1994. As regards Y_i, PI_i, and DI_i, the work is the same as in Mélitz and Zumer (1998), which contains a detailed description (in an appendix). The CANSIM database also provides consumption for residents by province and current spending by each Provincial government, the sum of which, divided by provincial population, yields C_i. The ten provinces in the study include all of Canada but the Yukon and Northwest Territories.

Italy

The main source of Italian regional data is "Conti economici regionali delle amministrazioni pubbliche e delle famiglie”, ISTAT, *anni 1983-1992, Argomenti*, n. 5-1996. Complimentary data come from the ISTAT computer database "Banco Dati Regio, Conti Regionali 1980-93”, available on diskette from ISTAT.

These data sources, however, do not allow for the regional decomposition of three elements in ASY's accounting: the consumption of lower-level governments, the net transfers from the central government to lower-level governments, and indirect taxes. More specifically, these data sources provide Y_i but only permit constructing C_i, DI_i and PI_i with respect to households and the net tax payments (net of transfers) of households to the central government. In other words, the hard work of Asdrubali, Sørensen and Yosha for the US in combining lower-level governments with households and decomposing indirect taxes in defining C_i, DI_i and PI_i would need to be repeated for Italy in order to apply the identical statistical approach to that country. Only for Canada does this work come ready-made in the data set.

With this caveat in mind, the ISTAT "Conti economici regionali delle amministrazioni pubbliche e delle famiglie" provides the right series for household disposable income DI_i (reddito lordo disponibile delle famiglie). To pass from DI_i to PI_i, it is simply necessary to add the personal income tax (imposte correnti), personal contributions to social insurance (contributi sociali effettivi e figurativi) and to deduct social transfers to individuals (prestazioni sociali nette). All the required data is available for 1983-1992. The rest of the needed regional data is found in the computer data base, and concerns population (popolazione residente a meta anno), gross domestic product (prodotto interno lordo per abitante), and private consumption (consumi finali interni per abitante).

United Kingdom

The previous qualifications with respect to Italy concern the UK as well. Subject to these caveats, the essential series are available on diskette directly from the Office for National Statistics (formerly the Central Statistical Office). As regards PI_i, it is necessary to begin with the data for "total personal income", which is exclusive of personal income taxes, but includes transfer payments. Transfer payments, or "social security benefits and other current grants from the general government", must then be deducted to obtain PI_i. Finding DI_i requires starting from "total personal income" once again and then deducting "personal income taxes" and "personal contributions for social security." The consumption series come from the table "Regional Accounts: Consumers' expenditure by Standard Statistical Regions", and the GDP series from the table "Regional Accounts: Gross Domestic Product by Industry." All of the series are available for 1971-1996 inclusively.

OECD and European Union

The data are available for 1970-1994 from the OECD *Economic Outlook electronic database*, except for GNP, which comes from the OECD *National Accounts* volumes. All the series, with the exception of GNP, can be found in real terms (at 1990 prices). As regards GNP, we used the GDP deflator in the preceding electronic database in order to obtain real values.

TABLE 1
THE ASDRUBALI-SØRENSEN-YOSHA (ASY) MODEL
USA 1964-1990

	(1) ASY	**(2) Replication**	**(3) Pooling without constraint**	**(4) Pooling with constraint**
b_K	0.39	0.34	0.41	0.34
t	(13)	(26)	(30)	(26)
$\bar{R}^2$	–	0.55	0.41	0.33
b_G	0.13	0.10	0.11	0.10
t	(13)	(17)	(17)	(17)
$\bar{R}^2$	–	0.77	0.18	0.19
b_C	0.23	0.26	0.33	0.18
t	(4)	(9)	(11)	(13)
$\bar{R}^2$	–	0.24	0.08	0.06
b_U	0.25	0.29	0.16	0.39
t	(4)	–	–	
$\bar{R}^2$	–	–	–	

Column 1 : ASY estimates : ASY(1996), Table I.
Column 2 : Replication with our econometric program.
Column 3 : Same as column 2 with use of ratios instead of time dummies.
Column 4 : Same as column 3 with constraint $1 - \beta_K - \beta_G - \beta_C = 0.39$.

TABLE 2
THE ASY MODEL WITH AVERAGE DATA
USA 1964-1990

	(1) k = 1	**(2)** k = 3	**(3)** k = 5	**(4)** k = 10	**(5)** "Between"	**(6)** "Between" with constraint
b_K	0.39	0.44	0.36	0.47	0.43	0.55
t	(13)	(21)	(12)	(47)	(42)	(56)
$\bar{R}^2$	—	—	—	—	0.56	0.56
b_G	0.13	0.16	0.17	0.18	0.04	0.11
t	(13)	(16)	(18)	(18)	(8.4)	(21)
$\bar{R}^2$	—	—	—	—	0.05	0.05
b_C	0.23	0.07	0.05	– 0.17	– 0.26	0.03
t	(4)	(1.1)	(0.6)	(– 0.6)	(– 16)	(3.1)
$\bar{R}^2$	—	—	—	—	0.16	0.16
b_U	0.25	0.30	0.42	0.53	0.79	0.31
t	(4)	(4)	(5)	(18)	—	
$\bar{R}^2$	—	—	—	—	—	

Column 1 : ASY (1996), Tables I and IV.
Columns 2, 3, 4 : ASY model with first differences over k years between observations, where k = 3, 5 and 10 respectively. See ASY (1996), Table IV.
Column 5 : Our replication of ASY model (with our definitions of variables and our econometric program) with averages of annual first differences over the entire study period: i.e., " between " estimates.
Column 6 : Same as column 5 with constraint $1 - \beta_K - \beta_G - \beta_C = 0.31$.

TABLE 3
POOLING ESTIMATE OF THE REVISED ASY MODEL
WITH b_U = 0.39
USA 1964-1990

Eq. 1	$\beta_K = 0.24$ (7.6)	$\gamma_K(z) = -0.067$ (– 5.06)	$\gamma_K(n) = -0.008$ (– 0.64)	$\gamma_K(P) = 0.346$ (10.20)	$\gamma_K(r) = 0.009$ (4.19)	$\bar{R}^2 = 0.47$
Eq. 2	$\beta_G = 0.13$ (7.8)	$\gamma_G(z) = -0.006$ (– 0.90)	$\gamma_G(n) = 0.030$ (4.89)	$\gamma_G(P) = 0.031$ (1.82)	$\gamma_G(r) = -0.005$ (– 4.25)	$\bar{R}^2 = 0.20$
Eq. 3	$\beta_C = 0.24$ (6.8)	$\gamma_C(z) = 0.073$ (4.97)	$\gamma_C(n) = -0.023$ (– 1.63)	$\gamma_C(P) = -0.377$ (– 10.04)	$\gamma_C(r) = -0.005$ (– 1.87)	$\bar{R}^2 = 0.11$

TABLE 4
"BETWEEN" ESTIMATES OF THE ASY MODEL
USA 1964-1990

	The simple model with b_U = 0.31		The revised model with b_U = 0.31		
Eq. 1	β_K = 0.55 (56)	$\bar{R}^2$ = 0.56	β_K = 0.45 (37)	$\gamma_K(n)$ = – 0.09 (– 13)	$\bar{R}^2$ = 0.60
Eq. 2	β_G = 0.11 (21)	$\bar{R}^2$ = 0.05	β_G = 0.09 (13)	$\gamma_G(n)$ = – 0.02 (– 5.6)	$\bar{R}^2$ = 0.06
Eq. 3	β_C = 0.03 (3.1)	$\bar{R}^2$ = 0.16	β_C = 0.16 (13)	$\gamma_C(n)$= 0.12 (16)	$\bar{R}^2$ = 0.04

TABLE 5
SOME PRELIMINARY STATISTICS REGARDING CANADA, THE UK AND ITALY AS COMPARED WITH THE US

Country	number of regions	Number of years in the sample	β_U in the full sample of annual observations	coefficient of variance *
Canada	10	33 (62–94)	0.32	0.27
UK	11	25 (72–96)	0.61	0.10
Italy	20	9 (84–92)	0.55	0.25
US	50	27 (64–90)	0.39	0.29

* The coefficient of variance is the average over all the years in the sample of the interregional variance of output per capita divided by the national output per capita.

TABLE 6
THE SIMPLE ASY MODEL FOR CANADA, THE UK AND ITALY

ESTIMATION METHOD	CANADA 1962-1994	UK 1972-1996	ITALY 1984-1992
Pooling	$\beta_U = 0.37$	$\beta_U = 0.61$	$\beta_U = 0.55$
b_K **t** $\bar{R}^2$	0.30 (8.60) 0.25	0.34 (8.74) 0.29	0.49 (9.33) 0.46
b_G **t** $\bar{R}^2$	0.08 (5.75) 0.16	0.0004 (0.017) 0.007	– 0.005 (– 0.19) 0.14
b_C **t** $\bar{R}^2$	0.25 (7.43) 0.28	0.047 (1.11) 0.12	– 0.039 (– 0.74) 0.10
Between	$\beta_U = 0.33$	$\beta_U = 0.60$	$\beta_U = 0.56$
b_K **t** $\bar{R}^2$	0.27 (26.1) 0.53	0.11 (4.72) 0.10	0.23 (3.52) 0.13
b_G **t** $\bar{R}^2$	0.004 (0.44) 0.12	0.09 (6.52) 0.16	0.13 (4.66) 0.23
b_C **t** $\bar{R}^2$	0.40 (33.4) 0.49	0.20 (10.11) 0.14	0.09 (1.31) 0.14

TABLE 7

POOLING ESTIMATES OF THE REVISED ASY MODEL ADDING SELF-FINANCING ONLY

	CANADA 1962-1994 b_U = **0.37**	UK 1972-1996 b_U = **0.61**	ITALY 1984-1992 b_U = **0.55**
Eq. 1	β_K = 0.23 (5.79) $\gamma_K(z)$ = – 0.014 (– 3.53) $\bar{R}^2$ = 0.28	β_K = 0.19 (0.97) $\gamma_K(z)$ = – 0.05 (– 0.73) $\bar{R}^2$ = 0.29	β_K = 0.65 (3.96) $\gamma_K(z)$ = 0.05 (1.02) $\bar{R}^2$ = 0.46
Eq. 2	β_G = 0.10 (6.20) $\gamma_G(z)$ = 0.004 (2.5) $\bar{R}^2$ = 0.17	β_G = – 0.18 (– 1.40) $\gamma_G(z)$ = – 0.07 (– 1.43) $\bar{R}^2$ = 0.012	β_G = – 0.13 (– 1.77) $\gamma_G(z)$ = – 0.04 (– 1.78) $\bar{R}^2$ = 0.009
Eq. 3	β_C = 0.30 (7.65) γ_C (z) = 0.010 (2.5) $\bar{R}^2$ = 0.28	β_C = 0.37 (1.67) γ_C (z) = 0.12 (1.48) $\bar{R}^2$ = 0.06	β_C = – 0.07 (– 0.44) γ_C (z) = – 0.01 (– 0.22) $\bar{R}^2$ = 0.10

TABLE 8

POOLING ESTIMATE OF THE COMPLETE REVISED ASY MODEL WITH b_U = 0.37 CANADA 1962-1994

Eq. 1	β_K = 0.17 (3.07)	$\gamma_K(z)$ = – 0.011 (– 5.46)	$\gamma_K(n)$ = – 0.066 (– 1.39)	$\gamma_K(P)$ = 0.766 (3.60)	$\gamma_K(r)$ = 0.008 (0.78)	$\bar{R}^2$ = 0.35
Eq. 2	β_G = 0.09 (3.77)	$\gamma_G(z)$ = 0.002 (1.77)	$\gamma_G(n)$ = – 0.045 (– 2.21)	$\gamma_G(P)$ = 0.054 (0.58)	$\gamma_G(r)$ = – 0.019 (– 4.37)	$\bar{R}^2$ = 0.24
Eq. 3	β_C = 0.38 (6.94)	$\gamma_C(z)$ = 0.010 (4.68)	$\gamma_C(n)$ = 0.111 (2.33)	$\gamma_C(P)$ = – 0.819 (– 3.84)	$\gamma_C(r)$ = 0.011 (1.11)	$\bar{R}^2$ = 0.29

TABLE 9
COEFFICIENTS OF VARIANCE OF GROSS DOMESTIC PRODUCT INTERNATIONALLY

	Unadjusted*	Corrected**
OCDE 23	1.59	1.13
UE 15	1.18	0.99

* The coefficient of variance is the average over all the years in the sample of the international variance of output per capita divided by the international output per capita.

** The difference between the unadjusted and the corrected values is that even though the national series are already in real terms (converted into dollars), we have subtracted the variance of the real effective exchange rate from the unadjusted coefficient of variance in making the correction. The corrected values are then more comparable to the earlier regional ones in the last column of Table 5.

TABLE 10
POOLING ESTIMATES OF THE REVISED ASY MODEL BASED ON INTERNATIONAL EVIDENCE

Estimation method	**OCDE 23 1960-1994**	**EU 15 1960-1994**
Pooling	$b_U = 0.80$	$b_U = 0.77$
Eq. 1	$b_{K1} = 0.05$ (10.23) $g_{K1}(z) = -0.001$ (–0.56) $\bar{R}^2 = 0.01$	$b_{K1} = 0.08$ (4.77) $g_{K1}(z) = 0.008$ (0.91) $\bar{R}^2 = 0.01$
Eq. 2	$b_C = 0.01$ (1.49) $g_C(z) = 0.001$ (2.38) $\bar{R}^2 = 0.01$	$b_C = 0.02$ (0.71) $g_C(z) = -0.015$ (– 0.97) $\bar{R}^2 = 0.03$
Eq. 3	$b_{K2} = 0.13$ (17.57) $g_{K2}(z) = -0.008$ (–2.45) $\bar{R}^2 = 0.06$	$b_{K2} = 0.13$ (6.22) $g_{K2}(z) = 0.006$ (0.51) $\bar{R}^2 = 0.02$
Between	$b_U = 0.80$	$b_U = 0.77$
Eq. 1	$b_{K1} = 0.07$ (21.21) $g_{K1}(n) = 0.006$ (2.71) $\bar{R}^2 = 0.03$	$b_{K1} = 0.13$ (14.55) $g_{K1}(n) = 0.059$ (8.76) $\bar{R}^2 = 0.10$
Eq. 2	$b_C = 0.05$ (8.12) $g_C(n) = -0.013$ (–2.95) $\bar{R}^2 = 0.01$	$b_C = 0.01$ (0.96) $g_C(n) = -0.068$ (–7.58) $\bar{R}^2 = 0.10$
Eq. 3	$b_{K2} = 0.08$ (13.26) $g_{K2}(n) = 0.007$ (2.10) $\bar{R}^2 = 0.14$	$b_{K2} = 0.09$ (9.55) $g_{K2}(n) = 0.010$ (1.42) $\bar{R}^2 = 0.02$